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Is There An Intra-Household Kuznets Curve?

Lawrence Haddad
and
Ravi Kanbur

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Is there a "Kuznets curve" for intra-household inequality? Does intra-household inequality first increase, peak, and then decrease as the household becomes better off?

Haddad and Kanbur found both theoretical and tentative empirical support for this hypothesis.

The policy significance of this finding is that the benefits of an increase in household well-being need not fully "trickle down" to the most disadvantaged members of the household. This is particularly true for the poorest households.

This finding should be taken into account in the design of supplementary feeding programs, for example. Research is now under way on this topic.

Is There An Intra-Household Kuznets Curve?
Some Evidence from the Philippines

by
Lawrence Haddad
and
Ravi Kanbur*

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* The authors thank Frances Halstead and David Neilson for their help with the graphics.

1. Introduction

The notion of a link between inequality and average well-being at the level of an economy was first introduced by Kuznets (1955). Relying on long-run historical data, and on a model of the development process, Kuznets hypothesised what has now become known as the "Inverse-U" relationship: as average well-being increases, inequality within a nation first increases, reaches a peak, and then decreases. A vast literature has grown up around this topic and there have been many attempts to test the relationship for data on currently developing countries (see Ahluwalia, 1976, Anand and Kanbur, 1990).

The Kuznets relationship was hypothesised for inequality in the economy as a whole. In recent years, however, much has been written about inequality within a household. As evidence on patterns of intra-household allocation has mounted (see Sen, 1984; Harriss, 1986; Deaton, 1989; Behrman, 1990) it has been recognised that an increase in the average well being of a household need not necessarily feed through uniformly to all members of that household. This raises the issue of what does happen to intra-household distribution of well-being when the household as a whole gets better off. It also raises the intriguing question: Might there exist a Kuznets-type inverse-U curve for intra-household inequality?

The question is indeed intriguing, and curiosity might well be sufficient reason for investigating it further. However, the investigation is in principle also relevant to the design of poverty alleviation policies. Is the best way of helping poor individuals to help poor households? Put another way, if we ensure an increase in the total consumption of poor households, are we in fact also ensuring an increase in the consumption of the most disadvantaged members of society? These will be recognised as being the micro-counterparts to questions on the "trickle-down" hypothesis at the macro level. But now there is a sharper policy issue. Targeting to individual, disadvantaged members of a household may be extremely difficult and costly. Are these costs worth paying? The answer depends on whether household level benefits do indeed "trickle down" to the individual level. Hence the importance of the intriguing question of what happens to intra-household distribution when the household as a whole becomes better off.

The plan of the paper is as follows. Section 2 reviews the theoretical arguments for the possible existence of an intra-household Kuznets-curve. We find theoretical support for the possible existence of the relationship. Section 3 then moves to an empirical investigation, with an attempt to look for the relationship in a data set on nutritional intake in the Philippines. Section 4 concludes the paper by indicating areas for further research.

2. Theoretical Arguments for the Inverse-U

Much of the literature on intrahousehold allocation has proceeded in the framework of household welfare maximisation (see for example Pitt and Rosenzweig, 1985 and Behrman, 1988a, 1988b). Let the household welfare function be given by

$$(1) \quad \underline{W} = W(x^1, L^1, x^2, L^2 \ldots\ldots x^n, L^n)$$

where x^i is a vector of consumption by the i^{th} individual, and L^i is labour supply. If we suppose that labour supplied is itself a function of consumption[1], the household's problem becomes

(2) Max W

$$\text{s.t. } p^x \sum_{i=1}^{n} x^i = \sum_{i=1}^{n} w^i L^i(x^i)$$

where p^x is the is the price vector for x, and w^i is the wage rate for individual i.

The solution to (2) determines the i. .rahousehold allocation, but does not, at this level of generality, generate useful insights on the relationship between intrahousehold inequality and average household well-being. Stiglitz (1976) specialised equation (2) into a form that could yield such insights. Let there be a single good, no labour supply term in the welfare function, and let $L^i(x)$ be identical for all i. If we further simplify to a separable Utilitarian welfare function for a "family farm" household with two identical members, the problem in (2) becomes

(3) $\text{Max } U(x^1) + U(x^2)$

$$\text{s.t. } x^1 + x^2 = G(L(x^1) + L(x^2);\ \theta)$$

where U(.) is an increasing concave function, $G(.,\theta)$ is the family farm's production function, and θ is a parameter representing labour productivity; $G_\theta > 0$.

1 Some evidence of this is presented in Haddad and Bouis (1989).

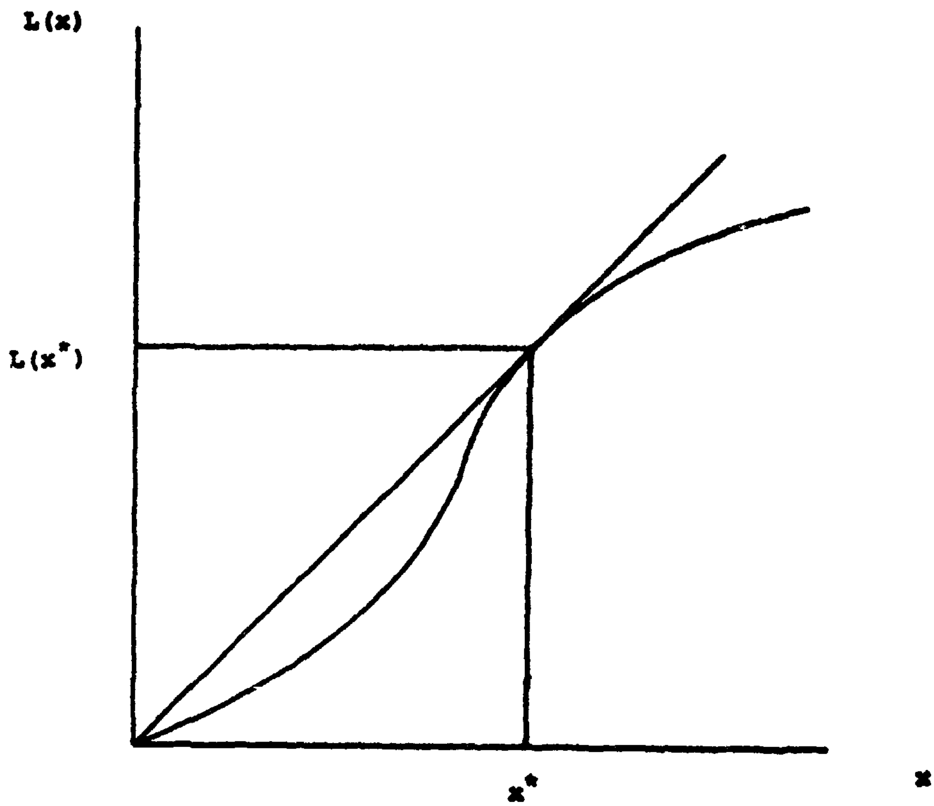
Stiglitz's (1976) analysis is based on a shape for the L(x) curve that has an increasing returns portion, as is common in the "efficiency wage" models. This is shown in figure 1. On this figure, x^* is the efficiency consumption - that which maximises labour supply per unit of consumption. If the value of θ was θ^*, such that

$$(4)\quad x^* + x^* = G(L(x^*) + L(x^*);\ \theta^*)$$

then clearly (x^*, x^*) is the solution to problem (3). In other words, for the household for which $\theta = \theta^*$, there will always be equality. *However*, when $\theta < \theta^*$ the household is operating in the *convex* portion of the L(x) curve. Now there is a tradeoff between equality in consumption and the total level of labour supply - since more can be generated by disequalising consumption. For $\theta \geq \theta^*$, the household is in the concave portion of the L(x) curve, so that disequalising consumption will not increase total labour supply. Thus equality preference in the concave shape of U(.) will dominate and we will continue to have equality.

We see then that for $\theta \geq \theta^*$ there is equality, while for $\theta < \theta^*$ there may be inequality. These results are established formally by Stiglitz (1976). Another case he considers is where θ is so low that the marginal product of labour is zero. In this case, even though creation of consumption inequality may lead to greater labour supply, this does not bring about any increase in total output, so the household prefers equality. Stiglitz (1976) summarizes the results as follows:

Figure 1: A labour efficiency curve



"In effect, when the economy is very well off, there is no trade-off between efficiency and equity, so maximisation of family welfare involves complete equality... On the other hand, when the economy is very poor, i.e...., the marginal productivity of labour services is zero; then of course again there is no tradeoff between equity and efficiency" (Stiglitz, 1976).

In between these extremes we find inequality. What has just been described is, of course, none other than an inverse U-shape!

A number of writers have eschewed the common preference, household welfare maximisation approach to intrahousehold resource allocation,9 and have instead emphasised individual preferences, conflict, and bargaining among household members (Manser and Brown, 1980; McIlroy and Horney, 1978; Folbre, 1984, 1986). While some have pointed to the difficulties in distinguishing empirically between the two types of models (Thomas, 1989; Senauer et. al. 1988), there nevertheless appears to be good reason why a more conflictual approach to intrahousehold allocation should at least be investigated (how else, for example, do we explain the observation that consumption patterns within a household depend on <u>who</u> earns what income?). Within the framework of two-person cooperative and non-cooperative bargaining theory, Haddad and Kanbur (1990a) have investigated the relationship between intra-household inequality and various types of improvement in average household well-being. While no general conclusions can be drawn (as in the case of household welfare maximisation models) they do show that under certain conditions bargaining models *also* predict an inverse-U relationship between intrahousehold inequality and average household well-being, as the result of interactions between the effects of increases in the total resources being bargained over, and changes in bargaining strengths.

Theoretical argument does seem then, to present a *prima facie* case in favour of a Kuznets-type Inverse-U hypothesis for intra-household inequality. The task of the next section is to investigate its existence in the context of a particular dataset.

3. Evidence from the Philippines

3.1 The data and variables

This section uses survey data from the Philippines in an attempt to verify the existence of an inverted U-shape between intrahousehold inequality and average household well-being. Our measure of well being is calorie adequacy, which has been used and discussed in the literature.

The data are from a four-round stratified random survey of the predominantly rural southern Philippine province of Bukidnon. The survey was conducted over a sixteen month period in 1984-85 and investigates, among other things, the dietary intakes of 448 households, covering 2880 individuals (excluding those who are breastfeeding)[2]. Calorie adequacy is defined as calorie intake divided by calorie requirement. The following sub-sections take up a number of details regarding the construction of variables of interest.

2 For a much fuller account of the data set context, collection, and content see Bouis and Haddad (Forthcoming).

3.1.1 Measurement of energy intake

We use calorie intakes as calculated from 24 hour recalls by the mother, of food eaten by individual household members. This information will be subject to a number of biases and intra*individual* variation. The biases due to 24 hour intakes are well-known (USDA 1986a) and in an attempt to minimise the net effect of biases *and* make the dietary snapshot more typical, we use only four-round *average* recalls: a technique used for a number of years by the USDA for its National Food Consumption Surveys (USDA 1986b) and also by Behrman (1988a, 1988b).[3]

3.1.2 Measurement of energy requirement

It is not so much *intra* as *inter*individual variation that is the main problem in measuring individual energy requirements. Consider a group from a healthy population, with the same age, gender, weight, pregnancy/lactation status, in the same climate, with the same pattern and intensity of time use: interindividual variation in metabolic efficiency would lead to a normal distribution of energy requirements, even within this group. The mean of this distribution

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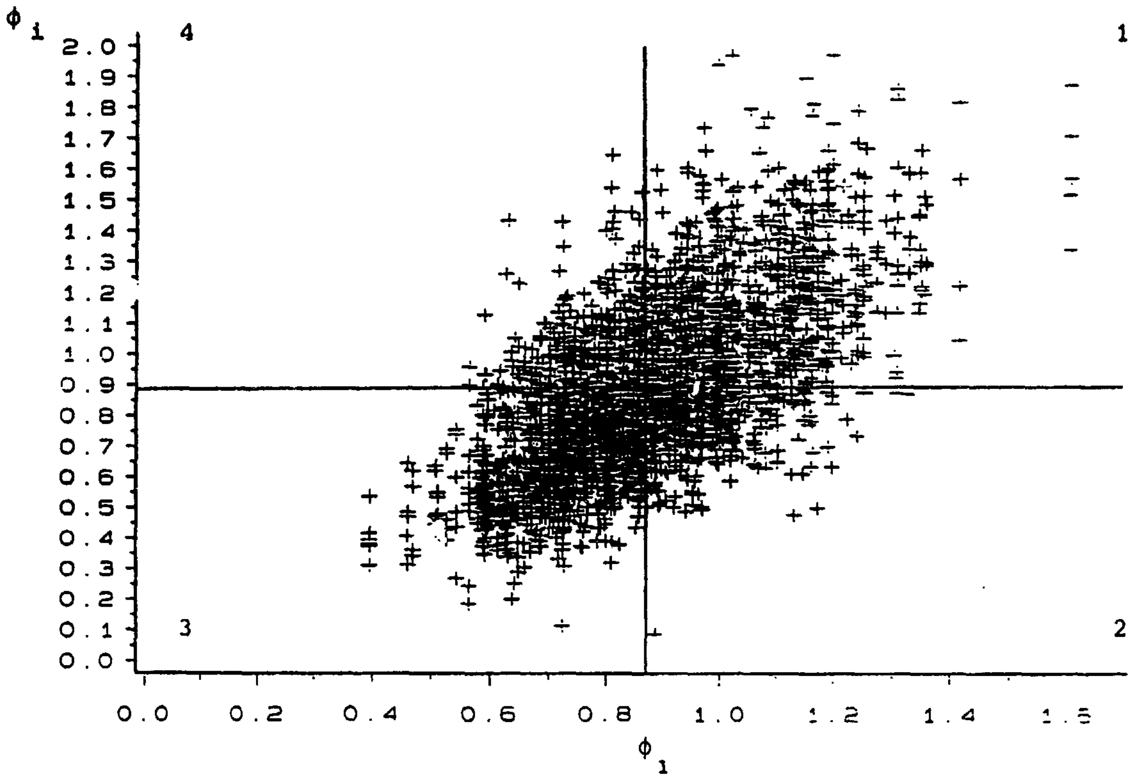
3 The position with respect to the 24 hour recall method is summed up by Chavez and Huenemann (in Sahn et. al. 1984): "One day may not represent a typical intake for the individual household. Twenty-four hour intakes of a large sample of households may, however, represent a typical daily intake for the community as a whole". Since total variation in individual energy intake can be represented as V(interindividual) + V(intraindividual)/n, we seek to minimise the intraindividual component by by averaging across four rounds of intake data (n = 4) for each individual. For each individual, each observation is independent, therefore we are able to reduce considerably the intraindividual component. Representativeness is further enhanced as data collection takes place across season and day of the week.

would be called the energy 'requirement' of an individual from this group. If the mean of the individual calorie intake/the group's energy requirement standard turned out to be 1.0, then half of the group would be classified as undernourished, although if the group were defined as homogeneously as possible with respect to requirement the *depth* of undernutrition would be minimised. Indeed, if the variance of requirements for age-sex different groups varied a great deal then calorie adequacy inequality within the household could be generated simply due to household composition effects. We attempt to control for household composition effects (behavioural or otherwise) with multivariate analysis, while calorie requirements (and therefore calorie adequacy) are based on Philippine age-gender-pregnancy status requirements (FNRI, 1976).[4]

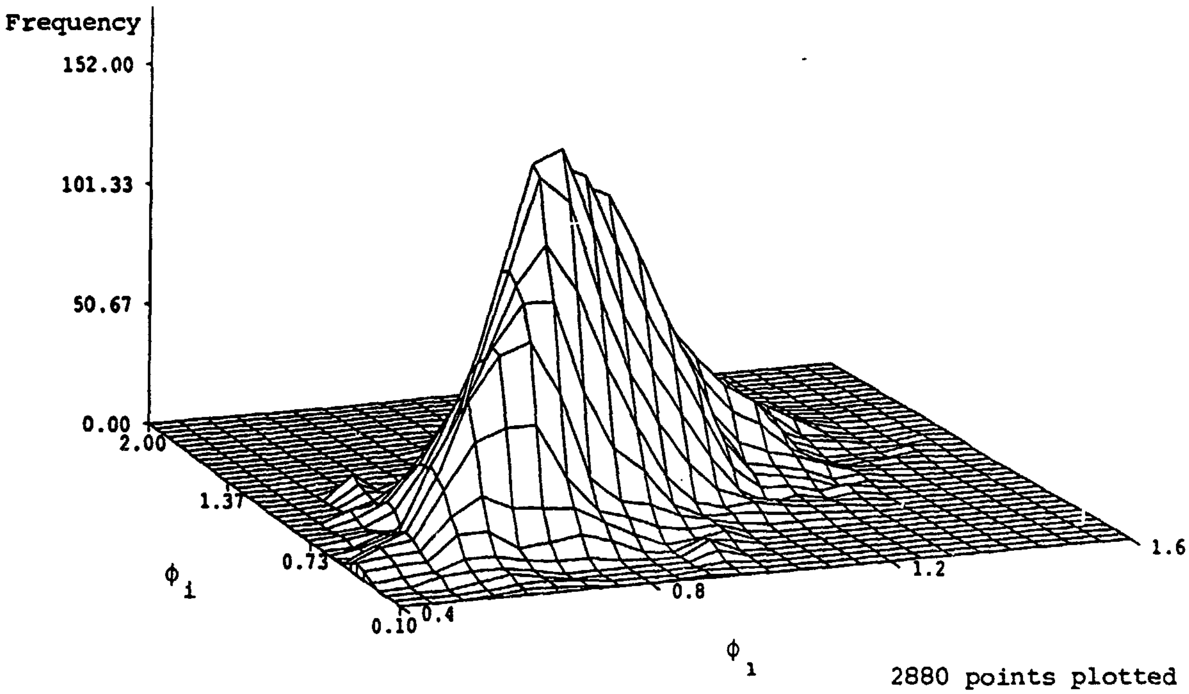
From the calorie intake and requirement data, a calorie adequacy variable, ϕ_i, is calculated for the ith individual and is used to construct ϕ_1 (the average ϕ_i within the household, or household welfare) and T_h, the Theil measure of inequality of ϕ_i within household h. Figure 2 presents scatter and relief plots of ϕ_i against ϕ_1. If all individuals were exactly meeting their requirements, the plots would collapse onto the (1,1) point. If there were different levels of household welfare, but everyone within a household had an equal value of ϕ_i ($=\phi_1$), the plots would represent straight lines at a 45 degree angle to the ϕ_1 axis.

4 A second set of calorie requirements was created by calculating basal metabolic rates for each individual by age, gender, reproductive status and body weight (see WHO 1985). The correlation coefficient between the first and second sets is high (0.878) and correspondingly the empirical results below differ little between sets one and two. A third set of estimates based on age-gender-pregnancy status-body weight-*activity pattern* requirements could not be constructed because of incomplete information on child activity patterns.

Figure 2: Plots of individual calorie adequacy (ϕ_i) against average calorie adequacy within the household (ϕ_1)



(a) scatterplot



(b) in relief

Points in quadrant 4 (2) represent individuals with calorie adequacies greater (lower) than the mean individual adequacy of 0.877, despite their location in households with mean calorie adequacies lower (greater) than 0.877. Forty-four percent of Figure 2 individuals from quadrant 2 are female with an average age of 9.36, while in quadrant 4 the corresponding figures are 51% and 31.93 years.

3.1.3 The dependent variable

The dispersion of ϕ_i within household h is captured by $T_h(\phi_i)$, the Theil measure of inequality (Kanbur 1984).[5] The Theil measure is given as:

$$T_h(\phi_i) = \sum_{i=1}^{nh} (1/n_h)[(\phi_i/\phi_1).\ln(\phi_i/\phi_1)].$$

For $n_h=1$, $\phi_i=\phi_1$ and $T_h=0$; but for $n > 1$, T_h is scale (ϕ_1) independent.[6] None of the 448 households have $n_h=1$. The dependent variable is logged as $\ln T_h$ since (i) u is truncated below $-x\beta$, and (ii) most of the values of T_h occur in a narrow band between 0.001 and 0.11 - logging the values expands the dispersion along the

5 The analysis presented below is little changed if we replace the Theil measure with another measure of inequality such as the log variance.

6 The Theil index is homogeneous of degree zero with respect to scale (see Kanbur 1984).

vertical axis. For our sample, Figure 3 plots T_h and $\ln T_h$ against ϕ_1.[7]

3.1.4 Right hand side variables

Apart from ϕ_1, our measure of household well-being, we include a standard set of explanatory variables (the reduced form equivalence of the competing underlying structural models, e.g. joint welfare or bargaining, is well known [Senauer et. al. 1988, Folbre 1986]). The explanatory variables include household demographics and the opportunity costs of time of the males and females in the household who are of working age.[8] Definitions and descriptive statistics of these variables are presented in Appendix 1.

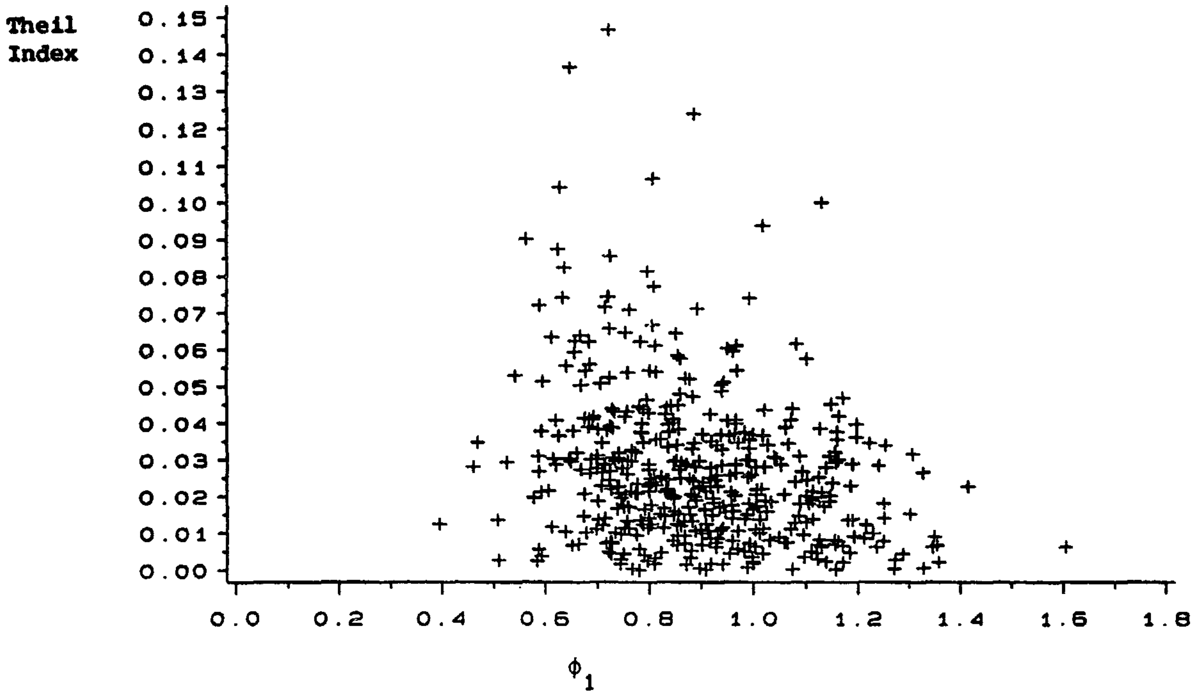
3.2 In search of the Inverse-U

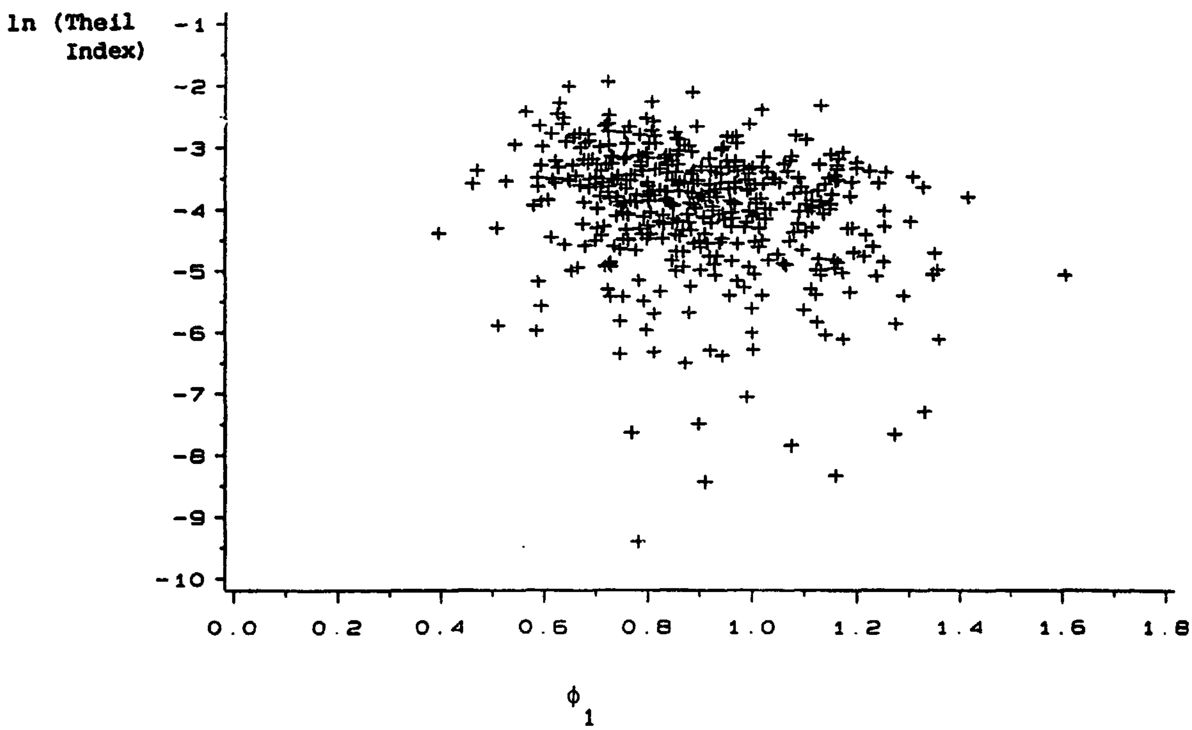
Rather than examine a whole array of non-nested functional forms that permit a non-constant $\delta T_h/\delta\phi_1$ across ϕ_1 values, we run a grid search for the spline cutoffs that minimise the residual sum of squares. The spline technique fits linear segments to the data, with the data determining the location of the cut-offs along the ϕ_1 axis, and is

7 In response to the usual worries about the mean as a measure of central tendency sensitive to extreme values, these figures were repeated for the median of ϕ_i within each household with the plots being very similar in both scale and pattern.

8 The opportunity cost of time is represented as predicted wage rate, calculated with the appropriate reservation-wage selectivity correction (Haddad 1987).

Figure 3: Scatterplots of the Theil index per household against average calorie adequacy within the household (ϕ_1)





448 points plotted.

generally regarded as less restrictive than functional forms that involve a transformation of ϕ_1.[9]

3.2.1 Linear spline analysis

Table 1 reports the estimated OLS coefficients on ϕ_1 in each of three line segments as significantly different from zero - and positive, negative, and negative respectively with the cutoffs at $\phi_1 = 0.68$ and 0.74.[10] The corresponding elasticities at ϕ_1=0.5, 0.7, and 1.0 are 1.780 (=3.599*0.5), -5.81 (=-8.304*0.7), and -0.53 (=-0.53*1.0). Figure 4, which plots the results in Table 1, seems to indicate that we have found an inverse U between $T_h(\phi_i)$ and ϕ_1, although the first and third segments are only significantly different from zero at the 7% and 9% levels respectively. The other explanatory variables have estimated coefficients as expected: controlling for ϕ_1, the demographic variables all increase inequality, with the preschooler and child variables having the largest effects - with no significant differences across gender. The adult male and female opportunity cost of time variables reduce inequality to similar extents although their estimated coefficients are barely larger than their estimated standard errors.

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9 The only restrictions imposed on the fitted curve are (i) the line segments are linear and (ii) consecutive segments meet at the boundaries. Clearly the importance of the former restriction diminishes as the number of segments increases (the segments are only linear in $\ln T_h$, ϕ_1 space anyway) and the importance of the second restriction can be tested against a model where the segments are disconnected (Stewart and Wallis 1987).

10 A crude grid search over the entire ϕ_1 space was conducted followed by a fine grid search in the neighbourhood of the first-stage minimum. The corresponding grid search for the second set of ϕ_i estimates produced cutoffs of 0.69, 0.79 and similiar slope coefficient estimates.

Table 1: Non-linearities in the relationship between calorie adequacy inequality (lnTh) and mean calorie adequacy within the household (ϕ_1): grid search for optimal selection of the two middle spline segments.

variable	estimated coefficient (OLS)	estimated standard error	t-stat	significance level
Constant	-4.864	0.595	-8.173	0.000
Z1	3.599	1.936	1.859	0.064
Z2	-8.304	3.153	-2.634	0.009
Z3	-0.533	0.319	-1.671	0.096
PWAGE1	-0.010	0.006	-1.725	0.085
PWAGE2	-0.016	0.015	-1.038	0.300
F15	0.367	0.078	4.707	0.000
M15	0.175	0.073	2.380	0.018
F614	0.132	0.041	3.186	0.002
M614	0.166	0.042	3.910	0.000
FGE15	0.082	0.069	1.187	0.236
MGE15	0.120	0.058	2.070	0.039

Adjusted R Square = 0.16840
F = 9.22910
n = 448

Notes:

1. line segment1=0.395 and line segment4=1.61, the extremes of the ϕ_1 distribution.

2. the regression with line segment2=0.68 and line segment3=0.74 is the optimal spline (see above):

   where,

   Z_1 = 0 if ϕ_1 lt 0.395
   Z_1 = ϕ1-0.395 if ϕ_1 ge 0.395 and lt 0.68
   Z_1 = 0.68-0.395 if ϕ_1 ge 0.68

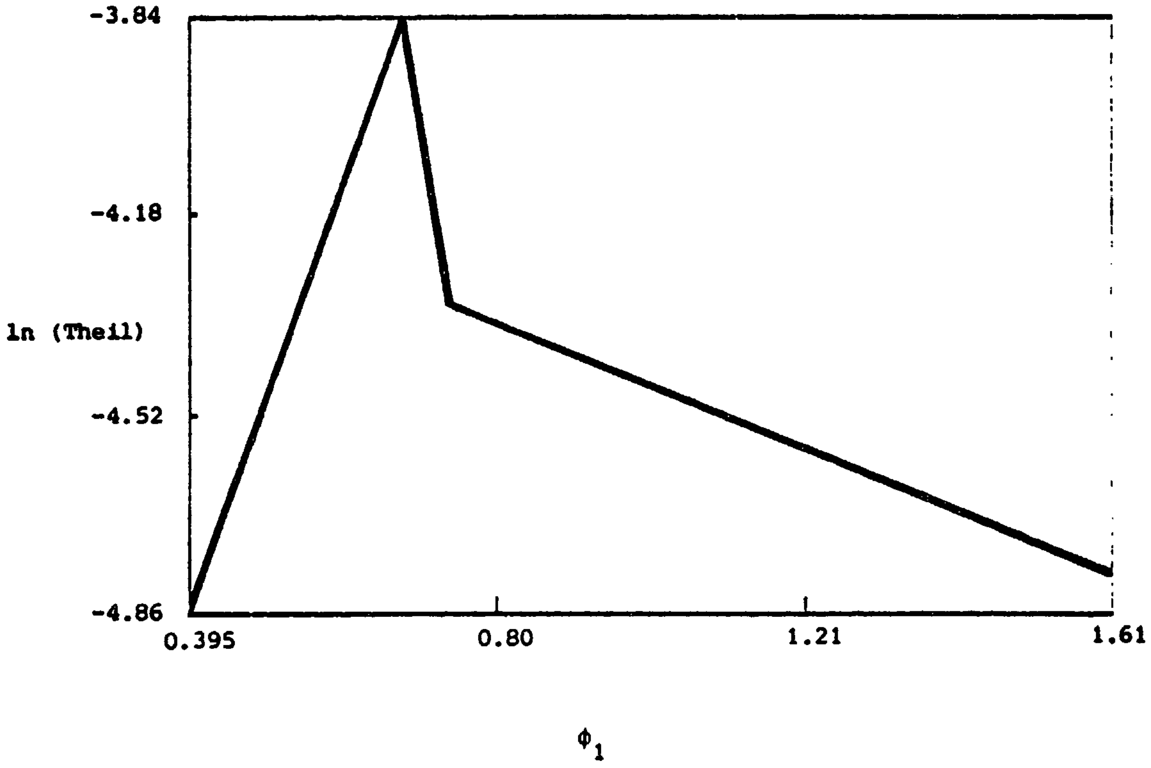
   Z_2 = 0 if ϕ_1 lt 0.68
   Z_2 = ϕ1-0.68 if ϕ_1 ge 0.68 and lt 0.74
   Z_2 = 0.74-0.68 if ϕ_1 ge 0.74

   Z_3 = 0 if ϕ_1 lt 0.74
   Z_3 = ϕ1-0.74 if ϕ_1 ge 0.74 and lt 1.61
   Z_3 = 1.61-0.74 if ϕ_1 ge 1.61

3. The households split into three groups of 49, 50, and 349 (sums to 448).

[d:\poverty\iii\table.2]

Figure 4: Spline fit of ln (Theil) against ϕ_1



3.2.2 Diagnostics on residuals

The results above are a first iteration awaiting diagnostic tests on the residuals. Does it matter from a statistical viewpoint that ϕ_i forms the basis of the dependent variable T_h *and* ϕ_1? Is the variance of the disturbance term positively related to household size (as is suggested by the above estimated coefficients on the household demographic variables)? How sensitive are the results to outliers in the (0.68, 0.74) interval for ϕ_1? Finally, how restrictive is the spline technique's insistence on line segments being connected?

- *simulteneity bias*

The potential for simulteneity bias is obvious as $\mathrm{cov}[\ln T_h(\phi_i), \phi_1(\phi_i)]$ is non-zero, but its *empirical* importance remains to be seen. Predictably, implementation of a standard Hausman (1978) test on the linear spline model rejected the null hypothesis of no simulteneity bias ($t = 2.107$) at 5% (but not at 1%). The three slope coefficient estimates over the range of ϕ_1 are inconsistent. Attempts to replace ϕ_1 with predicted ϕ_1[11], floundered due to the reduced range in ϕ_{1hat} (0.636, 1.197) compared to that for ϕ_1 (0.395, 1.610) - indicative of an important efficiency/consistency tradeoff. Apart from our implicit appeals to mean square error, our continued reliance on ordinary least squares is also based on the fact that we are more interested in the *sign* of the slope coefficients (is there an inverse-U?) rather

11 The instrument set consisted of spouse and wife heights and weights, price paid for corn and rice, and total area cultivated by household. The Hausman test result was robust to variation in the instrument set.

than the exact magnitude of the coefficients (how flat is the inverse-U?).

- *heteroscedasticity*

Since a household of two can, in principle, exhibit the same range of $\ln T_h(\phi_i)$ values as a household of ten, it seems reasonable to assume homoscedasticity of the disturbance with respect to household size. Nevertheless, the conditions under which the Theil measure is independent of household size are restrictive[12] and we test for heteroscedasticity with a White test (Maddala 1988) on the residuals of the linear spline function. The White test rejects the null hypothesis of homoscedasticity with respect to household size.[13] Weighted least squares estimates presented in Table 2 change little from the OLS estimates in Table 1.

- *outliers*

The outliers in Figure 3 raise the question of the sensitivity of our results to extreme values of $\ln T_h(\phi_i)$. Table 3 presents WLS residuals with absolute value greater than 3, together with their predicted WLS residuals and studentised WLS residuals (see Table 3 for

12 The impact on a household's Theil measure, of the introduction of a new member, depends on the redistribution that takes palce afterwards. If ϕ, remains unchanged, then T_h falls if the newcomer draws from initially better off members. If ϕ is allowed to change there are even more possibilities.

13 $\text{OLSresid}^2 = 4.904 - 1.141*\text{numhh} + 0.0691*\text{numhh}^2$
(t) (7.56) (6.55)

Heteroscedasticity does not affect the spline grid search which is based on OLS estimates [i.e. $RSS=\sum_h (T_h - X\beta_{hat})^2$].

Table 2: Non-linearities in the relationship between calorie adequacy inequality (lnTh) and mean calorie adequacy within the household (ϕ_1): WLS spline estimates

variable	estimated coefficient (WLS)	estimated standard error	t-stat	significance level
weighting factor	-4.796	0.497	-9.649	0.000
Z1	3.349	1.600	2.093	0.037
Z2	-8.142	2.413	-3.374	0.001
Z3	-0.448	0.309	-1.419	0.148
PWAGE1	-0.008	0.005	-1.816	0.070
PWAGE2	0.001	0.013	0.081	0.935
F15	0.169	0.067	2.539	0.011
M15	0.037	0.061	0.613	0.541
F614	0.077	0.037	2.092	0.037
M614	0.095	0.036	2.615	0.009
FGE15	0.119	0.051	2.340	0.020
MGE15	0.142	0.046	3.110	0.002

Adjusted R Square = 0.96089
F = 918.14447
n = 448

Note: weighted least squares regression use weights
= [(4.904-1.141* mhh+.0691*numhhsq)$^{-1/2}$]

[d:\poverty\iii\table.3]

Table 3: WLS residuals outlier analysis

hhid	wlsres	predicted	studentised	phil	lnTh	Th
276	-3.03	-3.32	-2.83	.75	-5.42	0.00442
357	-3.04	-3.06	-2.72	.78	-9.40	0.00008
356	3.19	3.31	2.90	1.02	-2.37	0.09340
288	-3.20	-3.34	-2.92	.98	-5.27	0.00514
311	-3.37	-3.49	-3.07	1.12	-5.82	0.00296
286	-3.60	-3.82	-3.32	.59	-5.57	0.00381
453	-3.76	-4.03	-3.49	.88	-5.67	0.00345
424	-4.25	-4.42	-3.90	.81	-6.31	0.00181

key:

hhid = household id

wlsres = weighted least squares residuals

predicted = predicted weighted least squares residuals [=lnTh-pred(lnTh) when the household's observation is dropped]

studentised = studentised weighted least squares residuals [=predicted wlsres/standard error]

phi_1 = $\phi 1$ for each household

lnT_h = ln of the Theil index for each household

T_h = the Theil index for each household

[d:\poverty\iii\table.4]

definitions). While the residuals are little changed in their predicted and studentised forms, 7 out of the 8 residuals are negative, implying that the third fitted segment may be being disproportionately dragged down by some extremely low values of T_h in the $\phi_1 > 0.78$ range, but this does little if anything to overstate the inverse-U. In addition, the six observations for which $T_h(\phi_i) >$ 0.1, were arbitrarily changed to $T_h(\phi_i) = 0.1$, with a reanalysis showing the numbers in Tables 1 and 2 changing little, and the conclusions changing not at all.

- *testing line segment continuity*

A simple F-test can be constructed testing the null hypothesis of continuous line segments [e.g. $\alpha_1 + \beta_1(\phi_1) = \alpha_2$; for segments 1 and 2]. Performing a joint test across the two kinks (number of restrictions = 2) we reject the null hypothesis ($F_{test} = 8.813 > F.^{01}{}_{2,418} = 4.61$) of continuity. Despite this rejection (and remember the alternative hypothesis constitutes free-fitted segments within each of the three ϕ_1 intervals) and the fact that coefficients on ϕ_1 within each segment are of reduced significance, they do display the same signs, and similar magnitudes to their spline counterparts. Indeed, as the results in Table 4 demonstrate, the spline technique actually understates the inverse U-shape.

4. Conclusions: Related Studies and Further Research

Following on from various <u>theoretical</u> possibilities of an intra-household Kuznets curve, we have conducted an <u>empirical</u> investigation for a particular dataset. The results, however tentative, provide

Table 4: Line continuity results

	spline		$\phi_1 \le .68$		$.68 < \phi_1 \le .74$		$\phi_1 > .74$	
Variable	ß	t	ß	t	ß	t	ß	t
constant	-4.864	-8.17	-6.964	-4.62	3.353	0.66	-3.755	-6.32
PHI1	-	-	3.611	1.79	-11.551	-1.61	-0.515	-1.48
Z1	3.599	1.86	-	-	-	-	-	-
Z2	-8.304	-2.63	-	-	-	-	-	-
Z3	-0.533	-1.67	-	-	-	-	-	-
F15	0.367	4.71	0.586	2.68	0.008	0.04	0.396	4.35
M15	0.175	2.38	0.618	2.35	0.304	1.58	0.134	1.60
F614	0.132	3.19	0.143	1.23	0.087	1.03	0.138	2.74
M614	0.166	3.91	0.131	1.19	0.142	1.34	0.154	3.06
FGE15	0.082	1.19	0.076	0.39	0.187	1.22	0.092	1.08
MGE15	0.120	2.07	0.357	2.23	0.244	1.55	0.081	1.19
PWAGE1	-0.010	-1.73	-0.053	-2.61	-0.013	-0.71	-0.008	-1.20
PWAGE2	-0.016	-1.04	0.047	1.23	0.027	0.66	-0.028	-1.57
Adj. R Sq.	0.168		0.265		0.037		0.140	
RSS	371.723		24.697		22.269		310.762	
n	448		49		50		349	

[d:\poverty\iii\Table.5]

preliminary support for a Kuznets curve at the very micro level of the household.

While we are unaware of any previous empirical studies that have tackled this question directly, we should mention here that the results of a number of studies can be interpreted within our framework. For rural Philippines, Senauer et. al. (1988) find that "the estimated wage rate of the wife and mother has a significant positive impact on the relative calorie allocation of both herself and her children and a negative effect on the husband's allocation", but their sample is too small to permit scale disaggregation. Folbre's (1984) analysis of the Philippine Laguna data shows that inequality of, in this case, work effort, was much less pronounced in wealthier families, although no comment is made about households on the lowest part of the wealth distribution.

Employing a similar technique to Senauer et. al., Haddad (1987) used the data set employed in this paper to report (i) that the calorie allocation of preschoolers relative to the household as a whole exhibited a U-shape across income quintiles, and (ii) mothers education and reported childcare time (instrumented) exerted the most positive and significant effects on relative preschooler calorie allocation within the middle income quintiles. These results are consistent with the inverted U-shape hypothesis we test here: - in the middle income ranges where increases in household income provide little comfort for the preschooler, non-monetary influences (such as mothers' education and time spent with child) become more important determinants of preschooler calorie allocation.

If the results reported here are confirmed with other datasets, then there appear to be significant policy implications for a strategy of reaching disadvantaged <u>individuals</u> through favouring disadvantaged <u>households</u>. It would appear that it is not simply enough to increase the total resources of a household since, particularly for poor households, the accompanying increase in inequality may well undermine the beneficial effects on the poorest individuals of the total resource increase. These effects should be taken into account in the design of supplementary feeding programes, for example, and research is now underway on this important topic (see Haddad and Kanbur, 1990b).

Appendix 1: Variable definitions and descriptives

variable label	mean	std dev	minimum	maximum	n	definition
TLNHH	.03	.02	.001	.15	448	Theil(T) per hh(ln)
LNTLNHH	-3.96	1.01	-9.40	-1.92	448	ln Theil(T)
PHI1	.89	.19	.395	1.61	448	mean of phi
M15	.59	.66	0.0	3.00	448	no. of males le 5 in hh
F15	.52	.62	0.0	3.00	448	no. of females le 5 in hh
M614	1.23	1.13	0.0	5.00	448	no. of males 5.01-14.99 yrs
F614	1.21	1.12	0.0	6.00	448	no. of females 5.01-14.99 yrs
MGE15	1.49	.90	0.0	6.00	448	no. of males ge 15 yrs
FGE15	1.39	.72	0.0	5.00	448	no. of females ge 15 yrs
PWAGE1	26.23	8.67	18.09	75.66	448	predicted daily wage, male hoh real pesos
PWAGE2	16.82	3.78	1.60	23.35	448	predicted daily wage, spouse real pesos
WT1	52.95	6.28	39.00	78.50	448	weight in kilos of male hoh
WT2	47.26	7.40	30.50	71.50	448	weight in kilos of spouse
HT1	161.18	5.86	145.30	184.00	448	height in cm, male hoh
HT2	150.16	6.35	101.30	166.40	448	weight in cm, spouse
CULTAREA	2.56	3.38	0.0	25.50	448	area cultivated, hectares
PRICEKG	5.48	.41	3.97	6.77	448	price/kg paid for rice
PCORNKG	4.26	.33	3.28	5.61	448	price/kg paid for corn
NUMHH	6.43	2.43	2.00	15.00	448	number in hh

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